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Monetary Transmission in an Emerging Targeter: The Case of Brazil

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Abstract

This paper lays out a structural model that incorporates key features of monetary transmission in typical emerging-market economies, including a bank-credit channel and the role of external debt accumulation on country risk premia and exchange rate dynamics. We use an SVAR representation of the model to study the monetary transmission in Brazil. We find that interest rate changes have swifter effects on output and inflation compared to advanced economies and that exchange rate dynamics plays a key role in this connection. Importantly, the response of inflation to monetary policy shocks has grown stronger and the output-inflation tradeoff improved since the introduction of inflation targeting.

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I. Introduction

The widespread adoption of inflation-targeting (IT) regimes by emerging-market (EM) economies has generated considerable interest in the channels through which monetary policy shocks affect output, inflation and other relevant aggregates. However, there has been a paucity of empirical research for EMs relative to the large literature on advanced countries, partly reflecting shorter time series and other problems not typically faced in studies of the latter.[2] The objective of this paper is to fill in some of this gap by developing a small macro model of the transmission mechanism of monetary policy in an economy with typical EM features and to empirically implement the model on data for Brazil. Brazil was one of the first EMs to formally adopt IT following a major currency crisis between late 1998 and January 1999 leading to the floating of the real, so that a longer time series under the new regime is available relative to other EMs.[3] Moreover, Brazil has consistently stuck to this monetary policy framework despite challenges created by some major external shocks and political gyrations. The central bank's response to inflation developments has often entailed high real interest rates (to levels among the highest in the world) and has generated considerable debate in policy circles. Thus, a formal quantitative assessment of the impact of such policy responses is clearly an issue of both analytical and practical interest to the general literature on IT.

We develop a small canonical macro model that is augmented to incorporate features known to be central to EMs. First, we allow for the role of external debt accumulation in driving the country risk premium and the exchange rate, relating these dynamics to exogenous fluctuations in the terms of trade as well as endogenous determinants of the external trade balance such as variation in domestic absorption. We show that the inclusion of an external debt accumulation equation in the model is not only of interest in its own right—as it permits the tracking of the effects of monetary shocks through key external aggregates—but it also imposes some stock-flow dynamics on the model that allows the model to have an invertible VAR representation. This allows us to estimate the model as a structural VAR (SVAR), where the effects of monetary policy shocks on output and inflation can be traced out through different channels.

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[2] Even though the literature on the relative performance of IT regimes in EMs is now sizeable (see, e.g., Loayza and Soto, 2002; Fraga, Goldfajn, and Minella, 2004; Miskin and Schmidt-Hebbel, 2007), model-based studies on the monetary transmission in these economies remain scarce. A notable exception has been the case of Chile. See Cespedes and Soto (2005) and the various references therein.

[3] Other EMs that subsequently adopted IT are: South Africa and Thailand in 2000; Korea, Hungary and Mexico in 2001; and Peru and the Philippines in 2002. Chile first introduced annual targets for inflation in 1991, but a full-fledged IT framework has been in place only since 1999 after the relaxation of capital controls and elimination of exchange rate bands.

We also augment the canonical model to include a bank-dependent domestic private sector. This allows us to capture additional effects of monetary policy shocks through the bank-intermediation channel emphasized by Bernanke and Blinder (1988), a channel which has been argued to be particularly relevant in emerging markets (Edwards and Végh, 1997). The effect of shocks to banks' lending capacity—arising, from say exogenous changes in reserve requirements and/or banking intermediation technology—on output, inflation, and other aggregates can also be traced out in the model.

We work with an SVAR representation of the model to identify the key features of the monetary policy transmission mechanism in Brazil and the speed they operate. The difficulty applying the SVAR strategy to EMs is the small sample of observations, implying that sensible theoretical restrictions become important for dimensionality reduction. This is particularly relevant for the model considered here because the unrestricted version of it has a large number of parameters. Imposing a factor structure may be a possible solution to the problem of making an SVAR parsimonious, but other strategies such as that followed in this paper should also be investigated. Achieving a balance between retaining a sufficient number of parameters to fit the data and the need for a relatively parsimonious specification to aid interpretation is often more art than science. Ultimately the acid test of such an exercise is in the results. As discussed below, our econometric estimates are consistent with both the prima facie evidence and theoretical priors. In particular, despite the relative short time span available for estimation (1999–2007), the SVAR estimates did not generate "price puzzles", "exchange rate puzzles" and other anomalies that abound in the literature using conventional unrestricted VARs.

The main results are as follows. First, compared with the United States and other advanced countries the transmission mechanism works much faster, with the bulk of the effects on output and inflation taking place within a year, which is consistent with the structural features of the Brazilian economy (e.g. shorter maturity of domestic credit and relatively high exchange rate pass-through) highlighted in the literature.[4] However, while the magnitude of the output gap and inflation responses to monetary policy shocks are smaller after the introduction of the IT framework in mid-1999, the effects on inflation are estimated to be larger, suggesting that the expectational channel may have become more important under IT. This is consistent with various institutional changes over the period, including stricter central bank adherence to the low inflation mandate and growing confidence in the IT regime.

Second, the exchange rate effects on inflation are found to be important even though Brazil is still a relatively closed economy to foreign trade. In particular, the exchange rate response to

[4] See Christiano, Eichenbaum and Evans (1998) for results based on data for the United States.

domestic policy interest rate changes is sizeable and akin to results from structural models that are based on uncovered interest parity.

Third, the bank credit channel is found to play a role in monetary transmission, although the net magnitudes involved are not very large. Our results are consistent with the existence of two channels through which monetary policy affects output via the bank credit channel, as highlighted in the literature (Kashyap and Stein, 1994; Edwards and Végh, 1997; Tornell and Westermann, 2005). On the one hand, by changing the lending-deposit spread, changes in the policy interest rate amplify the standard inter-temporal effect of monetary policy changes on absorption. On the other hand, there is an intra-temporal effect: monetary tightening tends to appreciate the exchange rate in the short run and this has expansionary effects on bank credit. The latter occurs because the domestic business sector tends to have a sizeable stock of foreign-currency denominated debt and the non-tradable sector of the economy is more "bank-dependent" than its tradable counterpart, implying that the overall demand for bank credit will tend to increase as relative prices shift towards non-tradable producers. This combination of balance-sheet effects of currency mismatches and the greater bank dependence of domestic firms—which have been shown to be typical of emerging markets (see, e.g., Tornell and Westermann, 2005)—imply that monetary policy will have non-trivial effects on bank lending and hence on absorption. Our estimates indicate that, while the inter-temporal channel eventually wins out, so that monetary tightening (loosening) depresses (boosts) bank credit, the intra-temporal channel appears to play an offsetting role.

One of the crucial issues arising in studies of the transmission mechanism using SVARs is the need to measure the permanent components of series in order to compute "output gaps." We employ the Beveridge-Nelson (BN) method of extracting permanent components which, as we discuss below, is more consistent with the spirit of our gap model and helps to overcome problems associated with output gap measures based on two-sided filters such as the Hodrick-Prescott filter. Nevertheless, we also show that our main results are robust to these two alternative measures of the output gap.

The remainder of the paper is divided into seven sections. Section II reviews the existing evidence on the monetary transmission mechanism in Brazil and provides a motivation for the model and results. Section III lays out a structural model that motivates the SVAR used. The connection between the models of sections II and III is detailed in Section IV. Section V provides a discussion of alternative measures of the output gap. Section VI describes the data, which is followed by the presentation of the estimation results in Section VII. The paper concludes with a brief summary and discussion of the main findings.

II. Existing Evidence on Brazil

The introduction of the IT framework in Brazil in 1999 has generated significant interest in understanding the monetary transmission mechanism and, as a result, there has been a growing subsequent literature seeking to identify and measure the channels through which the central bank's policy interest rate (SELIC) affects output and inflation. A description of some channels and a discussion of the central bank of Brazil's model are provided in Bogdansky, Tombini, and Werlang (2000).

Minella (2003) estimates a four-variable VAR using the overnight interest rate, inflation, output and M1 over the period 1975–2000, breaking the estimation into three sub-samples: the “moderate” inflation period of 1975–85; the high inflation period (1985–1994); and the post-1994 low-inflation regime. He finds that inflation inertia declines in the post-1994 period and that there is only weak evidence that monetary policy affects inflation in this post-stabilization period, even though his estimates point to significant effects of monetary policy shocks on output. Minella notes that this may be because of an identification problem because the sample (1994-2000) included a period dominated by interest rate responses to financial crises and defense of the exchange rate peg. A possible reason for this “anomalous” result is that the exchange rate was not included in the VAR.

The role of the exchange rate as a determinant of Brazilian inflation has been acknowledged in other studies. Bevilaqua, Mesquita and Minella (2007) find that the large appreciation of the real since 2005 has significantly contributed to the fall in inflation. In a similar vein, Favero and Giavazzi (2004) find that exchange rate movements affect inflation expectations, and, through this channel, the central bank interest rate setting. This suggests that the exchange rate may not only affect current inflation by changing the cost of imported goods, but there may be an important expectational channel at work as well.

Some attention has also been devoted to the interest rate reaction function. Using an HP-filtered measure of the output gap, Minella and others (2003) find that the parameter on the output gap in the monetary policy reaction function has the wrong sign and is not statistically significant from zero. They argue that this could arise because of simultaneity bias caused by supply shocks that depress the output gap and raise inflation. In theory supply shocks should be netted out of measures of the output gap, and any failure to do so will induce misspecification errors. The same study uses a smaller VAR estimated on monthly data including the pre-IT period (1994–2002) and finds that exchange rate volatility has been an important source of inflation variability in Brazil.

This paper expands on the above literature in a number of ways. First, we augment the set of equations used in previous studies to capture important open-economy linkages that were ignored in previous work. This involves separating out absorption and output effects as well as introducing an external liability equation. Aside from the obvious macroeconomic importance of current account deficits, we show that introducing changes in net external liabilities to the model ensures invertibility and hence a VAR representation.

Second we incorporate an equation for private bank credit growth to the SVAR. Indeed, much of the credit in the Brazilian financial system is still short term and firms are largely dependent on domestic banks for funding, having limited access to international capital markets. This is another channel that has not attracted enough attention in recent studies of the monetary transmission mechanism.

Third, we attempt to take some account of previous findings in Favero and Giavazzi (2004) about the role of external (EMBI) spreads in determining exchange rate risk, the fiscal stance, and hence output and inflation expectations. This is done by including the real US money market rate (which studies show is an important determinant of country spreads) and terms of trade in our SVAR. Since the two variables are exogenous for a small open economy, they do not need to enter as additional equations in the system, thus conserving on degrees of freedom.[5]

Fourth, we discuss the problems associated with measuring potential output with the Hodrick-Prescott (HP) filter and instead utilize the Beveridge-Nelson (BN) decomposition to obtain our estimates. We show that it produces more sensible results and is more consistent with the time series properties of the Brazilian data.

Finally, we restrict the sample to the IT period as a response to evidence that there have been large structural changes around the point of its introduction. For instance, Tombini and Lago Alves (2006) present recursive estimates of “free market” inflation showing that there has been a marked change in both inflation inertia and in exchange rate pass-through since 1999. In section VII we examine the issue of coefficient stability and ask if the response of inflation to monetary policy shocks has strengthened over time in the IT period.

III. The Structural Model

Our starting point is a small canonical macro model that has been used with some variants in the macroeconomics literature and has often been deployed for analysis at the IMF and by various central banks. It is derived from Euler-based equations on consumption and investment (which we implicitly aggregate into domestic absorption), a Phillips-curve type of equation for inflation and a Taylor-type rule relating the policy-controlled interest rate to expected inflation and the output gap. The model can be written as

[5] Technically the SVAR is an SVARX system due to the presence of exogenous foreign variables.

$$\tilde{n}_t = h_{lead}E_t\tilde{n}_{t+1} + h_{lag}\tilde{n}_{t-1} + h_r\hat{r}_{t-1} + \varepsilon_t^n \tag{1}$$

$$\tilde{y}_t = \beta_r\hat{z}_{t-1} + \beta_f\tilde{y}_t^* + \delta\tilde{n}_t + \varepsilon_t^y \tag{2}$$

$$\hat{\pi}_t = \alpha_{ld}E_t\left(\hat{\pi}_{t+1}\right) + \left(1-\alpha_{ld}\right)\hat{\pi}_{t-1} + \alpha_y\tilde{y}_t - \alpha_z\Delta\tilde{z}_t + \varepsilon_t^\pi \tag{3}$$

$$\hat{i}_t = \gamma_{lag}\hat{i}_{t-1} + \left(1-\gamma_{lag}\right)\left(\gamma\pi E_t\hat{\pi}_{t+1} + \gamma_y\tilde{y}_t\right) + \varepsilon_t^i \tag{4}$$

$$\tilde{z}_t = \delta_z E_t\tilde{z}_{t+1} + \left(1-\delta_z\right)\tilde{z}_{t-1} + \left(r_t - \hat{r}_t - \theta\hat{d}_{t-1}\right) + \varepsilon_t^z \tag{5}$$

$$\Delta\hat{d}_t = \bar{d}\left[\left(1-\bar{\pi}\right)\hat{i}_{c,t} - \left(1+\bar{i}_c\right)\hat{\pi}_t\right] + \omega_n\tilde{n}_t - \tilde{y}_t + \left(\omega_m - \omega_x\right)\hat{z}_t - \omega_x\hat{tt}_t \tag{6}$$

$$\hat{i}_{c,t} = \hat{i}_t^* + \theta\hat{d}_{t-1}, \tag{7}$$

where a levels deviation will be designated by a ^ while the log-level deviations are given a ~.

The first equation provides a specification for the log of domestic absorption (n_t), which is simply GDP minus net exports.[6] It is measured as a log deviation from some "equilibrium" value and so should be regarded as a "gap" variable. Models that emphasize gaps are a convenient way of organizing policy and forecast discussions so that one can concentrate separately upon where one sees the system heading and the path of adjustment to that point. Most modern macroeconomic models can be written as gap models and so the approach is fairly flexible. The equilibrium value may be a constant or time varying.[7] In this case the absorption gap depends upon the real rate of interest $\hat{r}_t$, also measured as a deviation from some equilibrium value. The definition of the real rate will involve an expected inflation rate, which for simplicity will be proxied with the inflation target. Notice that no other variable determines the level of absorption, which is consistent with the standard Euler equation for

[6] By domestic absorption (N) we mean the sum of expenditures on consumption, investment and government purchases, so by definition GDP will be equal to N plus exports minus imports.

[7] The need for the latter often reflects the fact that there are permanent stochastic components that need to be removed to induce stationarity in the measured gap. Later we will provide a discussion of such transformations in detail, but even if there were no permanent components in the data, it is often the case for emerging-market economies that the equilibrium values to which the system will be adjusting are shifting over time in response to structural changes in the economy. Consequently, care should be taken when constructing these gaps, and any assessments of the resulting measures should rely on institutional knowledge of the economy being studied. This knowledge can be quite informative, not only of the presence of structural changes (such as those in the transition from high to low inflation regimes and across monetary policy frameworks as in Brazil between 1998 and 1999), but also of how sensible one's estimation results appear to be. A striking example in the empirical macro literature of the problems arising from ignoring country-specific features in broad cross-country regressions pertains to the identification of the long-run effects of fiscal deficits on inflation; although solidly backed by theory, these effects are not easily discernable without properly taking into account country--group specific features in the estimation strategy. See Catão and Terrones (2005).

consumption with habit persistence in DSGE models.[8] Implicit in such a specification is that the other variables making up absorption, principally investment, are also functions of a real interest rate. While it might be worth considering augmenting this equation with some expressions for the rate of return on investment and other measures of the actual relative price or cost of capital (including tax wedges for instance), such measures are not readily available for emerging markets, and Brazil is no exception.

The second equation describes an output gap, connecting the logs of GDP$\left(y_t\right)$, domestic absorption (*n*), imports and exports. Since imports are determined by total expenditure and the real exchange rate, while exports are linked to the real exchange rate and foreign expenditure, we simply eliminate imports and exports to produce a relation linking the logs of GDP and domestic absorption, the log of the real exchange rate z_t and the log of foreign output y_t^*. The specification assumes that there is a 1 quarter lag between the real exchange rate and trade flows, which is quite common in empirical specifications because of the presence of lags between orders and deliveries. We also add a shock to the equation to allow for the fact that, in many emerging-market countries, the opening up of the economy produces an import surge that is much larger than expected from the price and output elasticities for import demand. Although, much of this movement can be accounted for by a time-varying equilibrium value for the import share, one will probably want some of the changes observed to be captured by a shock that is persistent.

The third equation provides a specification for inflation that includes the output gap and exchange rate. As suggested by previous studies and confirmed by the data reported in Figure 1, the exchange rate has a very significant role to play in influencing the price of tradables and CPI inflation. Following other studies inflation (π_t) is expressed as a deviation from the target level of inflation.

The next equation is a monetary policy reaction function, where $\hat{\imath}_t$ is defined as the nominal interest rate less the target inflation rate. The parameter $\gamma \mathrm{lag}$ seeks to capture the degree of interest rate smoothing in central bank policy, which is usually highly significant in policy reaction functions and Brazil has been no exception to this in recent years (See Figure 1 as well as the econometric evidence in Section VII).

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[8] As discussed in the next section, we will augment this canonical specification to include the role of domestic interest spreads - a wedge which arises in models with deposit- and credit-in-advance constraints (see, e.g., Edwards and Végh, 1997). Since the domestic interest spread is itself a function of the policy interest rate as well as of a measure of the supply side of bank credit, this baseline specification for absorption will remain unchanged except for the addition of an extra term using an "excess credit" measure.

The exchange rate equation is a flexible version of UIP, with the latter holding if $\delta_z = 1$. Note that the exchange rate is defined such that a rise in it represents an appreciation. There is also a "risk premium" in the equation, determining the level of the domestic interest rate relative to the foreign rate in equilibrium. As standard in many real business cycles models of small open economies (see, e.g., Schmitt-Grohe and Uribe, 2003), the risk premium is a function of the level of foreign debt relative to GDP. This is not only appealing from a theoretical perspective, but it is also consistent with recent external developments in many emerging markets (including Brazil), where a decline in net external debt has been accompanied by a decline in standard measures of country risk, such as the EMBI spread (See Figure 3).

Finally, the last equation is a log-linearized version of the identity describing how the level of debt to GDP changes over time, the derivation of which is provided in the Appendix. As shown below, the consistency between the debt stock and flows which this imparts to the model is important in making an invertible VAR possible. Expressing it as a log-linearization around long-run equilibrium is consistent with the overall spirit of gap models, and is also a standard way of closing open economy real business cycle models (see Schmitt-Grohe and Uribe, 2003). In this equation, the terms of trade $\left(\hat{tt}\right)$ are taken as exogenously given - a reasonable assumption for a small open economy; $\omega_{m,}\omega_x$ and ω_n are the import, export and absorption shares in GDP, respectively; $\bar{d}$ is the steady-state level of net foreign debt and $\hat{\psi}$ is nominal GDP growth. Note that since the evolution of the terms of trade and the trade balance determine the path of net external debt through this identity, and the level of debt affects the exchange rate through the risk premium term, those variables will also be potentially important determinants of the exchange rate, output and inflation.

IV. SVAR Representation

The above system represents a conventional view of the monetary policy transmission mechanism. If expectations are ignored it has a triangular structure, thereby creating the potential for the system to be written as an ordered SVAR. To develop this point further consider the absorption equation, which is

$$\tilde{n}_t = h_{lead} E_t \tilde{n}_{t+1} + h_{lag} \tilde{n}_{t-1} + h_r \hat{r}_{t-1} + \varepsilon_t^n$$

Now, if ξ_{t-1} are the lagged endogenous variables and x_t the exogenous variables, we might define the information available at time *t* as ξ_{t-1} and x_t. Then $E_t\left(\tilde{n}_{t-1}\right) = \xi_{t-1}^{'} f_1 + x_t^{'} f_2$ leading to

$$\tilde{n}_t = h_{lead} \xi_{t-1}^{'} f1 + h_{lead} x_t^{'} f_2 + h_{lag} \tilde{n}_{t-1} + h_r \hat{r}_{t-1} + \varepsilon_t^n ,$$

producing an SVAR(1) equation (actually an SVARX(1) system). To get a higher-order SVAR it is necessary to assume that the shocks have higher-order serial dependence. In particular, if ε_t^n is an AR(1) then the equation will have second order lags in it.

One possible way to get the above model closer to an SVAR is to augment each of the structural equations with extra lags in the variables and to then estimate that system. However, in practice this was not feasible given the large number of parameters. Moreover, one feels that the standard assumption in DSGE models that shocks are known when expectations are formed is somewhat limiting, but if we discard such an assumption one cannot easily use standard programs for estimation. For this reason we follow the standard SVAR methodology and do not make a distinction between expectations and structural dynamics. In any event, since the existing time series on inflation expectations in Brazil as well as in other emerging are usually very short (only available for the post-2002 period in the case of Brazil), this also poses clear limitations to empirical estimation of models that strictly separate dynamics and expectations for these countries.

In general it will be the case that every variable affects the conditional expectation, but simulations of the Berg, Karam and Laxton (2006) model (which is a variant of the canonical model above), using the calibrated values they provide, suggest that many of the variables are of little significance. If enough data was available empirically this could be investigated by regressing $\tilde{n}_{t+1}$ against ξ_{t-1} and x_t and making an assessment of which variables are important in forming the conditional expectation. Given the sample sizes we will be working with there are obvious limitations to the strategy. Instead we start with the idea that a minimal set of determinants of the expected value of a variable, when it is the dependent variable of some structural equation, would be the regressors of that equation and their lags. Some limited experimentation might then be done to see if there are extra variables that are useful in predicting the variable whose expected value needs to be found. To illustrate this consider the absorption equation. We know that if its shock is an AR(1) then the solution will have the form

$$\tilde{n}_t = \phi \tilde{n}_{t-1} + \sum_{j=0}^{\infty} \gamma^j E_t\left(\hat{r}_{t-1+j}\right) + \varepsilon_t^n$$

Now the expected value of the real interest rate can be investigated by regressing $\hat{r}_t$ against variables anticipated to determine the real interest rate. These would come from the Taylor rule and that points to lagged values of the GDP gap and inflation as being its determinants. Indeed a regression of $\hat{r}_t$ against two lags of itself and two lags of the absorption gap and inflation (proxying for the GDP gap) gives an R^2 of .90. Hence, using two lags of variables in the absorption equation and the inflation rate does a good job of capturing the expectations in an unrestricted way. In some cases it might make sense to form an estimate of the expected value of the real interest rate and to then use that in the equation but, given that the only new variable introduced into the absorption equation is the exchange rate, imposing such a restriction would mean little is gained. The same principles follow for the other equations in which expectations appear.

A second modification to the system above is to decompose the real interest rate into an unrestricted combination of the nominal rate less the target inflation rate and the deviation of inflation from target. This separation has two functions - it allows for expected inflation to deviate from the target and it also encompasses the possibility that the short-run effects of a change in nominal interest rates on expenditures might not be limited to their effects on real rates. There are many rigidities in the economic system that result in a given change in the nominal interest rate having a stronger effect in the short run than an equivalent change in the real rate would have - e.g. the front-end loading of many loans means that real repayments rise sharply as nominal interest rates change. This leads us to work with nominal interest rates and inflation—both measured relative to the inflation target—rather than combining them together in the form of a real interest rate variable defined as the nominal interest rate minus current inflation.[9]

A third modification is to allow for the role of the exchange rate in the monetary policy reaction function. To the extent that exchange rate changes help determine expected future inflation, and this determines the interest rate settings, we should proxy the expected inflation term by current and lagged inflation but also lagged exchange rates. This does not imply that the central bank is engaged in any exchange rate targeting but is simply a way of capturing expectations within an SVAR structure, which is consistent with the arguments and analytical framework laid out in Svensson (2000). The real exchange rate is used to keep the system compact, but also because price stickiness tends to mean that movements in the nominal exchange rate will be reflected strongly in $\tilde{z}_t$. Even if the inflationary effects are ignored, there is an alternative rationale that does suggest that exchange rates might appear in an interest rate rule. This comes from the fact that in an emerging market context there may well be "fear of floating" considerations whenever liability dollarization is extensive and welfare losses associated with sharp exchange rate movements are non-trivial (see, e.g., Calvo and Reinhart, 2002). Such outcomes might justify some "leaning against the wind" through interest rate policy.[10]

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[9] To the extent that target inflation is fully credible and thus becomes a true measure of expected inflation, then the interest rate variable in this regression is the theoretically relevant measure of the real interest rate in the Fisherian sense. This is consistent with econometric results reported in section VII.

[10] Another rationale for this specification arises when the central bank seeks to prevent or minimize excessive appreciation of the currency for external competitiveness reasons. Avoiding large appreciations in the disinflation process has been, for instance, a well-known feature of the Chilean experience under IT. This has

(continued…)

A fourth modification to the canonical model pertains, as discussed above, to the inclusion of a credit channel involving the amount of credit being granted by banks that is in excess of some "normal" level, pc_t. The channel envisages that "excess" credit would have effects on output and inflation over and above the standard monetary mechanism. Microfounded models of the credit channel featuring deposit- and credit-in-advance constraints as well as costly banking (Edwards and Végh, 1997; Goodfriend and McCallum, 2007) imply that a wedge appears in the Euler equation governing absorption, so that absorption becomes a function of lending and deposit spreads. Similar emphasis on the amplification mechanism associated with bank interest rate spreads is found in the earlier literature on the "credit view" (Bernanke and Blinder, 1988; Kashyap and Stein, 1994), also suggesting lending-deposit spreads ought to feature in an absorption equation like (8). As shown in Edwards and Végh (1997) for a typical emerging market context, such spreads are a direct positive function of the policy interest rate itself plus a term related to the credit-to-expenditure ratio. The latter seeks to capture, *inter alia*, the effects of shocks to bank technology as well as changes in minimum reserve requirements to credit expansion, hence affecting absorption as well as other relevant macro aggregates.[11] Thus, a parsimonious way of introducing such a credit channel into our model, so as to be consistent with both a micro-founded theoretical framework and the empirically-oriented literature on advanced economies (see Kashyap and Stein, 1994 for related references), is to add the ratio of the nominal level of credit to nominal GDP into the absorption equation - a variable which we will refer to as "excess credit". The "normal" level of credit is then implicitly made a function of standard determinants such as the base interest rate and absorption.[12]

Since at any point in time domestic bank credit is endogenously determined within a system, it is necessary to decide how to account for this within an SVAR. In order to allow for contemporary effects of credit expansion on expenditure, we place the excess credit variable before any other variable of the system. To be consistent with theory, we restrict the

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been reiterated recently by Chile's central bank governor who in a public statement on October 17, 2007 indicated that the central bank reserves itself the right to intervene in the exchange rate market if the real exchange rate exhibits an "important misalignment" (sic) with respect to fundamentals. See http: //ttda.today.reuters.com/news/Article.aspx?type=business News&StoryID=2007-10-18T1780731Z_01_N18424725

[11] The specific way in which Edwards and Végh (1997) model bank technology yields a relationship between the lending and deposit spreads (measured relative to the base interest rate) and the credit to deposit ratio. But since deposit-in-advance constraints imply that deposits are proportional to expenditure, this directly translates into a functional relationship between bank spreads and the credit to expenditure ratio.

[12] Allowing for the presence of an autonomous component in the "excess credit" variable that is not directly related to the interest rate seems particularly appropriate in the case of Brazil. Indeed, the existence of a large development bank (BNDES), which accounts for up to one quarter of domestic credit, and whose lending policies and rates arguably respond to other incentives can result in some lending rates significantly below market rates.

explanatory variables entering the excess credit equation to be absorption, the real interest rate (separated into its nominal and inflation components), and the real exchange rate. The latter can enter the equation for two reasons. First, there may be sizeable balance-sheet effects of the type documented in Calvo, Izquierdo and Mejia (2004).[13] Secondly, it is possible that the non-tradable sector may be more "bank-dependent" than the tradable sector, as suggested by evidence on some emerging markets (Tornell and Westermann, 2005, ch.6). Both structural features imply that a real exchange rate (REER) appreciation will increase real credit demand in the non-tradeable sector, leading to higher aggregate credit. Although Brazil is not as dollarized as many other emerging markets, significant balance-sheet effects may also be present, and they could further strengthen the positive impact of a REER on domestic credit. As illustrated in Figure 4, there is indeed *prima facie* evidence from Brazilian data that is consistent with a positive association between the REER and domestic credit growth, often with changes in the real exchange rate being followed by changes in domestic credit growth. Section VII provides supportive econometric evidence that real exchange rate appreciations tend to foster domestic credit growth.

The strategy then is to embed the structural model in some SVAR of a chosen order.[14] It should be noted that, although we split the real rate into a nominal rate and an inflation component in empirical work, for simplicity we write the relations out in terms of a real rate when presenting the equations for the SVAR. The debt equation remains as in equation 6. One might replace this with an SVAR equation and, although the impulse responses are much the same quantitatively, they are much less smoother than when an identity is assumed. Moreover, treating it as an identity saves on the number of parameters to be estimated. Overall and without counting the debt equation, this set of structural restrictions in our model reduces the number of parameters to be estimated below an unrestricted VAR by fourteen in the case of a first-order SVAR and twenty-eight for a second-order one.[15]

$$pc_t = g_1 pc_{t-1} + g_3 n_{t-1} + g_6 \hat{i}_{t-1} + g_8 \hat{\pi}_{t-1} + g_9 \hat{z}_{t-1} \tag{8}$$

$$\tilde{n}_t = h_{1n}\tilde{n}_{t-1} + h_{1i}\hat{i}_{t-1} + h_{1\pi}\hat{\pi}_{t-1} + h_{lpc} pc_t + h_{lpc} pc_{t-1} + \varepsilon_t^n \tag{9}$$

[13] Even though the "dollarization" of private sector liabilities in Brazil is not nearly as extensive as in many other EMs, it is far from negligible. Starting from negligible amounts in the early 1990s, foreign currency denominated debt rose to 36% of total corporate debt in 1999, reaching 40% in 2002 (Bonomo, Martins, and Pinto, 2004, Table A.2). Using a large panel of firm-level data, Bonomo and others (2004) also find that balance sheet effects of currency movements have significant effects on credit demand and investment.

[14] As discussed further below, we have experimented with both SVAR(1) and SVAR(2) models, but opted for an SVAR(1) specification. With larger samples it might be desirable to work with a higher-order VAR.

[15] If the debt equation is also estimated, this would add an extra seven so treating it as an identity saves a number of parameters.

$$\tilde{y}_t = \beta_{ly}\tilde{y}_{t-1} + \beta_{lz}\hat{Z}_{t-1} + \beta_{0y^*}\tilde{y}^*_t + \beta_{1y^*}\tilde{y}^*_{t-1} + \beta_{0n}\tilde{n}_t + \beta_{1n}\tilde{n}_{t-1} + \beta_{1\pi}\hat{\pi}_{t-1} + \varepsilon^y_t \quad (10)$$

$$\hat{\pi}_t = \alpha_{l\pi}\hat{\pi}_{t-1} + \alpha_{0y}\tilde{y}_t + \alpha_{1y}\tilde{y}_{t-1} + \alpha_{1z}\tilde{z}_{t-1} + \varepsilon^\pi_t \quad (11)$$

$$\hat{i}_t = \gamma_{li}\hat{i}_{t-1} + \gamma_{0\pi}\hat{\pi}_t + \gamma_{l\pi}\hat{\pi}_{t-1} + \gamma_{0y}\tilde{y}_t + \gamma_{ly}\tilde{y}_{t-1} + \gamma_{lz}\tilde{z}_{t-1} + \varepsilon^i_t \quad (12)$$

$$\tilde{z}_t = \delta_{lz}\tilde{z}_{t-1} + \delta_{0r}\hat{i}_t + \delta_{lr}\hat{i}_{t-1} + \delta_{0r^*}\hat{r}^*_t + \delta_{lr^*}\hat{r}^*_{t-1} + \delta_{0\pi}\hat{\pi}_t + \delta_{l\pi}\hat{\pi}_{t-1} + \theta\hat{d}_{t-1} + \varepsilon^z_t \quad (13)$$

As discussed above and unlike many SVAR studies, we allow for a stock variable (external debt) to be present in the system. One may ask about the econometric benefits (or pitfalls) of having such a stock variable in our SVAR specification, i.e., is there any important difference between a system with just flow variables and one that also incorporates a stock of debt variable. To see this, consider a simplified model in which there is a desire to stabilize the level of debt relative to a target with some variable being manipulated to achieve that. If $\tilde{x}_t$ is the control variable and $\tilde{d}_t$ is the stock of debt defined as a gap relative to its desired equilibrium value, the identity relating to debt accumulation is taken to be

$$\Delta\tilde{d}_t = \tilde{x}_t,$$

where we assume a zero real rate of interest for simplicity.[16] In order to stabilize the debt we need a rule of the form

$$\tilde{x}_t = a\tilde{d}_{t-1} + c\tilde{y}_t + e_{t,} a < 0,$$

Where $\tilde{y}_t$ is an output gap. Then

$$\Delta\tilde{d}_t = a\tilde{d}_{t-1} + c\tilde{y}_t + e_t, \quad (14)$$

and the debt gap converges to zero since $\tilde{y}_t$ is a stationary process.

Now suppose we attempted to use an SVAR which did not include $\tilde{d}_t$ i.e. it only consisted of $\tilde{x}_t$ and $\tilde{y}_t$, as is common with many fiscal VARs and open economy studies. To see the effect of this we need to solve for $\tilde{d}_t$ and then substitute that variable out of the system. From (14) this gives:

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[16] In our case $\tilde{x}_t$ will be net imports.

$$\tilde{d}_t = \left(1-(1+a)L\right)^{-1}\left[c\tilde{y}_t + e_t\right]$$

so that the fiscal rule can be expressed as

$$\begin{aligned}\tilde{x}_t &= a\left(1-(1+a)L\right)^{-1}\left[c\tilde{y}_{t-1} + e_{t-1}\right] + c\tilde{y}_t + e_t \\ &= (1+a)\tilde{x}_{t-1} + ac\tilde{y}_{t-1} + ae_{t-1} + c\tilde{y}_t + e_t - (1+a)\left(c\tilde{y}_{t-1} + e_{t-1}\right) \\ &= (1+a)\tilde{x}_{t-1} + c\Delta\tilde{y}_t + \Delta e_t\end{aligned}$$

The error term Δe_t is a non-invertible MA(1), meaning that there is no SVAR representation for $\tilde{x}_t, \tilde{y}_t$ (there will be another equation for $\tilde{y}_t$ in the system but it may or may not involve the level of debt). Thus, in this case compression of the set of variables results in a VARMA process but, importantly, one in which the MA term is not invertible.

V. Producing Gap Measures

Many gap models measure gaps by an application of the Hodrick-Prescott (HP) filter. Generally, DSGE models derived from some economic optimization theory will be expressed in terms of the levels of variables and a transformation needs to be applied to express these in gap form. The transformation means that the gap model will have Euler equations with errors that consist of the original shocks along with terms involving the expected values of the change in the permanent components found with the HP filter—see Fukacs and Pagan (2006). Hence one needs to be careful when working with gap variables to examine whether the expected value of the change in permanent component distorts the shocks that are defined by the Euler equation.

As Harvey and Jaeger (1993) pointed out the HP filter can be regarded as extracting the signal P_t by applying the Kalman Filter to a state space model of the form:

$$\begin{aligned} z_t &= P_t + T_t \\ \Delta^2 P_t &= \nu_t \\ T_t &= u_t \\ \lambda &= \frac{\operatorname{var}(u_t)}{\operatorname{var}(\nu_t)}. \end{aligned}$$

The model clearly implies that

$$\begin{aligned} \Delta^2 z_t &= \Delta^2 P_t + \Delta^2 T_t \\ &= \nu_t + \Delta^2 u_t \end{aligned}$$

$$= e_t + \alpha_1 e_{t-1} + \alpha_2 e_{t-2}$$

where e_t is an uncorrelated process. Setting $\lambda = 1600$ we find that $\alpha_1 = -1.77, \alpha_2 = .8$. Fitting this model to Brazilian GDP data we get $\alpha_1 = -.75, \alpha_2 = -.25$. Of course this process has a common unit root to the MA and AR parts which cancels, implying that the log of Brazilian GDP is an *I*(1) process, which contrasts with the implied *I*(2) model when using a HP filter.

Because the HP permanent component is generated as an *I*(2) process it might be expected that this would lead to difficulties in using it in such models. Indeed in Laxton, Shoom and Tetlow (1992) it was found in a simulation experiment that, when the potential level of output actually followed an *I*(1) process, measuring the output gap by filtering the observed output with a HP filter, produced estimates for the parameter on the output gap in a Phillips curve that were well below the true values used in producing the simulated data.

This suggests that we want to utilize a measure of the permanent component of a series that is extracted under the assumption that it is *I*(1). One such filter that does this is the Beveridge-Nelson (BN) filter. Two methods are available for computing it. To contrast the assumptions with the HP filter we note that it can be viewed as extracting P_t from the following system using the Kalman filter—see Anderson, Low, and Snyder (2006)

$$z_t = P_t + \gamma \nu_t$$
$$P_t = P_{t-1} + \nu_t .$$

Alternatively, *P* can be computed as

$$z_t^{BN} = z_t + E_t \sum_{j=1}^{\infty} \Delta z_{t+j.}$$

Using the latter form it is clear that the output gap will be $-E_t \sum_{j=1}^{\infty} \Delta z_{t+j}$. If Δz_t is an AR(*p*) then $E_t \sum_{j=1}^{\infty} \Delta z_{t+j}$ will be a linear function of $\Delta z_t, \Delta z_{t-1}, ..., \Delta z_{t-p+1}$. Hence the BN measure of the output gap is constructed as the negative of an average of growth rates. Notice that this means that one will see a *negative* relation between the output gap and growth, so that regressing inflation against the growth in output should produce a negative coefficient on the latter.

It is sometimes said that the problem with the BN estimate of an output gap is that it is not smooth enough. If one uses a low order AR process to approximate Δy_t this may well be true but, when a higher-order AR is adopted, it is often much smoother (see, e.g. Morley, 2007). The intuition is that the gap is constructed by averaging growth rates, and that will generally result in some persistence in the output gap measure.

In applying the BN filter to the Brazilian data—necessitating the fitting of an AR—care should be taken owing to the fact that the moments in the output growth process appear to have changed dramatically in the post-1994 low inflation environment, as illustrated in Figure 2. So, both with BN and the HP filter estimation, we limit ourselves to the post-1994Q3 period, when inflation trends sharply downwards and growth becomes much more stable. Figure 2 plots the BN- and HP-filter measures of the real GDP gap with estimation beginning in 1994Q3. To avoid the end-point problem with the HP filter we extend the sample beyond 1997Q2 (the last date for which actual output is currently available) based on a projection of output growth to 2009 that has potential GDP eventually settling at a 4.5% annual growth rate, which is consistent with the latest projections in the IMF's World Economic Outlook. As can be seen from Figure 5, if one fits an AR(4) to Brazilian GDP data and extracts the BN estimate of the output gap from such a univariate process it resembles that found with the HP filter $(\lambda = 1600)$, though the latter is smoother overall.

VI. The Brazilian Data Set

We use quarterly data for the period 1999Q1 to 2007Q2. While consideration has been given to the use of monthly data, the unavailability of key national account aggregates on a monthly basis led us to work with quarterly data. The beginning of the sample (1999Q2) corresponds to the adoption of IT and so we are looking at the transmission mechanism as it would operate now rather than historically. Limiting the estimation to the post-1998 period is advisable in light of evidence from Tobini and Lago-Alves (2006) which is suggestive of significant structural changes in inflation dynamics and exchange rate pass-through in Brazil before and after 1999. Further, given the Brazilian triple-digit inflation experience prior to 1995, and the far-reaching changes in the price indexation system and inflation dynamics since that time (see Minella, 2003), it is certainly advisable to exclude the pre-1995 period at the very least.

Seasonally adjusted national income account data was taken from both the IMF's International Financial Statistics (IFS) and data provided by the Brazilian Planning Ministry Research Institute (IPEA). Domestic bank credit to the private sector was taken from the same sources and seasonally adjusted using the X11 routine in AREMOS. The real exchange rate series is from the IMF and is computed as a weighted average of exchange rates for nearly all trading partner using CPI deflators and 2000 weights. Foreign real GDP series is also from the IMF and is weighted by the trading partners' trade with Brazil. The inflation variable is seasonally adjusted CPI including both administered and “free market” prices. While it has been customary to separate the two based on the argument that “administered” prices have a stronger backward-looking adjustment component (largely due to the nature of the multi-year contracts between the government and the new incumbents in the utility industries privatized during the 1990s), we see this distinction as somewhat artificial. For a number of reasons it can be potentially misleading for the purposes of setting the monetary policy stance, and perhaps irrelevant if the task at hand is indeed to model aggregate inflation. A first reason is that administered prices still respond to demand pressures, even if with a longer lag, because of backward-looking indexation clauses in the underlying concessional contracts. Second, although utility prices are typically key inputs to "free market" prices, the interaction between the two is certainly complex so that, even though one might add both series into a VAR, this would be unlikely to address such complexity. Thirdly, the extent to which wage earners make a distinction between the types of inflation is unclear: if they only care about overall inflation, second-round effects will stem from this, and that would reduce the advantage of decoupling the two inflation rates. For these reasons, the estimation results reported below refer to the "all item" CPI. Corresponding estimates using the "free market" price component of the CPI were also performed, yielding essentially the same results although displaying, as expected, a slightly greater sensitivity to monetary policy shocks. These results are available from the authors upon request.

VII. SVAR Estimates

The SVAR described by equations 8-13 is fitted. After experimenting with a two-period lag length, we chose to focus upon an SVAR(1) as convergence to a constant debt level was ensured in all cases and the impulse responses were noticeably smoother. It is important to note that a VAR(1) in 7 variables implies that individual variables follow an ARMA(7,6) process, which is a very high order of dependence. Any higher-order is unlikely to be detectable from only 34 time-series observations. Gaps for absorption, GDP and foreign GDP are found by fitting univariate AR(4) processes and then extracting the Beveridge-Nelson transitory component. In parameterizing the debt equation ϖ_m, ϖ_x and ϖ_n were replaced by the average import, export and GNE ratios to GDP over 1995–2007, while $\bar{d}$ was set to the average net foreign debt/GDP data over the same period. Target inflation (4.5% in recent years) was used to compute $\bar{\pi}$.

There are clearly many parameters being fitted in each equation and, given that we have only 34 observations, one would expect that the individual t ratios of the coefficients are even less informative than they normally are for SVARs. Yet, individual equation estimates generally have signs conforming to theoretical priors and are statistically significant at (or close to) conventional significance levels. Their magnitudes are also broadly sensible, with single-equation R^2s that range from a low of 0.45 (inflation equation) to 0.90, most being around 0.80, as for the real exchange rate and excess credit equations. This can be seen from the point estimates for the three main behavioral equations in the model - those for inflation, the real exchange rate and the excess credit equation (t ratios in brackets)[17]:

$$\hat{\pi}_t = 0.31\hat{\pi}_{t-1} + 1.56\,\tilde{y}_{t-1} - 0.12\tilde{z}_{t-1}$$
$$(1.69) \quad (1.58) \quad (-2.16)$$

$$\tilde{z}_t = 0.62\tilde{z}_{t-1} + 0.69\left(\hat{i}_t - \hat{\pi}_t + \hat{r}_t^*\right) - 0.51d_{t-1}$$
$$(4.48) \quad (2.16) \quad (-2.44)$$

$$pc_t = 0.54pc_{t-1} - 0.08\left(\hat{i}_{t-1} - \hat{\pi}_{t-1}\right) + 0.08\tilde{z}_{t-1}$$
$$(3.63) \quad (-1.84) \quad (3.43)$$

Some other comments can be made about the nature of the estimated structural equations. The output gap enters the inflation equation with the right sign and a large coefficient (a one percentage point reduction in the gap lowers inflation by 1.5 percentage points) which is statistically significant at a level close to 10 percent. As noted above it is not infrequently the case that one obtains very low (and sometimes incorrectly-signed) and insignificant coefficients on the output gap in inflation equations. The inflation equation estimates also indicate that the exchange rate has a powerful effect on domestic inflation: a 10 percent appreciation in the real effective exchange rate (which took place for instance in Brazil between end-2006 and mid-2007) leads to a 1.2 percentage point reduction in inflation, *ceteris paribus.* Likewise, the exchange rate also has a non-trivial effect on domestic credit: a 10 percent appreciation leading to a drop in the ratio of credit to GDP close to 1 percentage point. These exchange rate effects are in turn driven by changes in the domestic real interest rate relative to the offshore real interest rate (in this case the CPI adjusted 3-month US-treasury bill) and net external debt, which enters the exchange rate equation with the postulated negative sign and a sizeable coefficient (-0.51), which is statistically significant at close to 1 percent.

Turning to the system-wide impact of shocks, while it is possible to identify a monetary shock in this SVAR, the absorption shock is hard to interpret, as it could come from either a transitory demand or supply shock. Since the GDP shock is found by conditioning upon absorption it represents change in preferences for domestic goods, which may either be from a rise in exports or a shift against importables. The inflation shock might be regarded as a

[17] Adding the GNE gap to the credit equation gives a coefficient of 0.38 with a t-ratio of 1.03. While yielding the right sign, this lowers the t-ratio on the real interest rate variable to 1.2 suggesting strong multicollinearity between the variables. This is consistent with the evidence that absorption is itself a function of the interest rate and lagged excess credit so the GNE gap is spanned once those two variables are already present in the equation. For the purpose of illustrating our point, we dropped the GNE gap from the above single equation estimates but do include it in the VAR.

supply side shock and the exchange rate shock is often interpreted as a shift in the risk premium. As indicated in the single-equation estimates above, $\hat{d}_{t-1}$ is highly significant in the real exchange rate equation (and thus represents how the risk premium changes with the level of external debt), which is a necessary restriction for the system to converge to a constant debt-GDP ratio.

The graphs below show the impulse responses (IRs) to the interest rate and bank credit shocks over the period 1999Q2–2007Q2. All variables are measured in percentages; inflation and interest rate are in annual equivalents of quarter-on-quarter changes. The shock to excess credit is for a 1 percentage point rise, which is about half of the standard deviation of that variable during 1999q2–2007q2. The interest rate shock is for a 100 basis point rise in the yearly interest rate. The adjustments on output and inflation are quite fast. Unlike many closed-economy models in which inflation takes longer to respond to interest rate movements than the output gap, in an open economy there is a separate channel operating through the exchange rate that can speed up the adjustment. Despite the evidence alluded to in section II of a declining pass-through, there is evidence that exchange rate effects on inflation remain strong in Brazil, as shown by the single equation estimates reported above.

The impulse responses in Figure 6 show that a positive shock to the interest rate appreciates the real exchange rate in the very short run. There is no "price puzzle" or "exchange rate puzzle" in the results. The effects of this shock on the GDP gap and inflation are not as strong as in the US for instance (see Christiano, Eichenbaum and Evans, 1998): in Brazil, the gap cumulatively declines by 23 bps in the first year and the annual inflation decline peaks at just under 40 bps in the third quarter after tightening; no less importantly, the bulk of the effects take place within four quarters. This entails a much shorter lag than in existing estimates for the US and the Euro area (see also Angeloni et al., 2003).

The difference between the impulse responses for absorption and output is clear and this is directly reflected in the trade gap dynamics, shown in the right bottom panel of Figure 6. This result indicates that the effects of monetary policy on expenditure last one to two quarters longer than on output. This is a similar result to what has been observed in other studies that make the distinction (Dungey and Pagan, 2000), where the argument was made that this can arise if monetary policy has a stronger effect upon expenditures that are relatively import intensive. Of course the initial exchange rate appreciation that follows from the positive onshore interest rate differential will itself tend to reduce the domestic component of expenditure relative to the import component; so, this latter effect tends to work in the opposite direction. Overall, the loss in competitiveness seems to win out on impact, with a shaper fall in GDP than in absorption.

The paths of real GDP and absorption also relate to the effects of monetary tightening on bank credit and external debt. Higher onshore-offshore interest differentials appreciate the

REER (via UIP) and boost external debt (e.g. through the carry-trade). Consistent with the underlying theory, the IRs suggest that the initial REER appreciation also tends to boost non-tradable sector demand for bank credit (through both higher relative price of non-tradables and a positive balance sheet effect), somewhat offsetting the negative effect of monetary tightening that works through the inter-temporal channel. This dampens somewhat the effects of monetary policy on absorption. Nevertheless, the intertemporal channel appears to dominate eventually, leading to a fall in excess credit and hence in absorption.

Turning to the shocks to bank credit, Figure 7 shows shock leads to a rise in absorption and, after one period, to a rise in inflation. The overall effects are not large. This may suggest that the "pure" credit channel is not particularly strong in Brazil, which is consistent with the fact that bank credit/GDP remains well below both advanced country levels and those observed in other emerging markets, notably Asia. But it may also mean that this ratio (and other related measures, such as bank spreads) are not entirely satisfactory proxies, especially in countries where a sizeable informal credit sector exist which is not captured in existing bank credit statistics. While beyond the scope of this paper, more work needs to be done to pin down the "pure credit" channel in such contexts, including through disaggregated firm or sector level evidence.[18] Inflation responses are less smooth and faster than is case for a monetary policy shock, consistent with hypothesis that bank credit - particularly where the average credit maturity is short - tends to speed up the monetary transmission mechanism.

Finally, we examine the robustness of these results in two ways. One is the choice of output and expenditure gap metrics. Figure 8 reports the IR results for the monetary tightening shock using the HP-filter measure of the output gap, again spanning the period 1999Q2 to 2007Q2. The IRs are similar to the one obtained with the BN filter. A main difference is the somewhat deeper fall in the output gap and the less damped response of absorption, translating into a more immediate positive impact on the trade balance. While the impact exchange rate appreciation is of similar magnitude as in the estimates using the BN filter (see Figure 6), the larger drop in the output and expenditure gaps imply that inflation declines by more, with a trough at 50 bps in the fourth quarter (vs. 40 bps using the BN filter). In contrast with BN filter-based IRs, the negative intertemporal effect on domestic credit is more immediate, i.e., it is not initially offset by the intra-temporal effect associated with the exchange rate appreciation, implying that domestic credit/GDP declines nearly monotonically and gradually after the positive interest rate shock.

Second, we look at the issue of whether the impact of monetary policy on inflation has grown stronger over the IT regime. This is plausible, not only because longer periods of low

[18] See the discussion by Eichenbaum (1994) on the difficulties faced by empirical work in the identification of a credit channel effects in the US, for which longer and better data and more disaggregated empirical evidence are available.

inflation arguably make agents more forward-looking, and therefore more responsive to the current monetary stance, but also if credibility in monetary policy management grows over time (see Goodfriend, 2005 for a further discussion of these arguments and references therein). We do this by simply re-estimating our SVAR(1) model further into the IT period. While we are clearly limited by the short length of our sample, Figures 6 and 7 report the IRF estimates for the period 2001q2–2007q2, i.e., two years after the onset of IT in Brazil. Clearly, while the responses are all in the same direction as before but less smooth - likely reflecting sample limitations - the overall effects of a 100 bps monetary tightening on output and inflation are stronger in the short run: the inflation decline now peaks at 70bps (as opposed to 35bps), while the GDP gap drops only 3 bps further, to 26 bps cumulatively in the first year after tightening.[19] This stronger response of inflation to monetary policy via the output gap can be seen by comparing the single equation estimates for the period 2001q2-2007q2 with those beginning in 1999q2 which we reported above:

$$\begin{aligned}\hat{\pi}_t = 0.41&\hat{\pi}_{t-1} + 2.18\tilde{y}_{t-1} - 0.11\hat{z}_{t-1}\\ &(2.07) \qquad (1.79) \quad (-1.70)\end{aligned}$$

A comparison of the estimated equations suggests that inflation has become more sensitive to the output gap (as gauged by the coefficient of 2.18 in the above regression compared to 1.56 for the period 1999q2–2007q2) and that the associated coefficient is also more precisely estimated, with t-ratio of 1.79 compared to 1.58 previously. Figure 11 plots how the coefficient on the output gap in the inflation equation has evolved over time when estimated recursively. Overall, the above results indicate that not only did inflation become more responsive to monetary policy, but also that the higher sensitivity of inflation to the output gap entails that a given rate of disinflation can be achieved with a smaller “sacrifice” in output relative to trend.

VIII. Conclusion

The adoption of IT by some EMs in the wake of the emerging market financial crises of the 1990s was initially met with some skepticism on the grounds that there was little (if any) evidence of well-established channels of monetary transmission in EMs similar to those in advanced economies.[20] Subsequent developments indicated that this view was off-the-mark:

[19] We have also experimented with extending the estimation period backwards to 1998Q2, i.e, after Brazilian inflation declined to single digit levels in 1997 and before the 1999Q1 devaluation and debt crisis. Both the coefficients on the output gap and inflation were significantly lower than after 1999Q2. This again reinforces the view that there was significant structural change in the monetary transmission process in Brazil before and after IT.

[20] See Mishkin (2004) for a survey of these views and Masson, Savastano, and Sharma (1997) for an early prediction of the inadequacies of an IT framework in EMs.

several EMs have since moved to IT regimes and succeeded in achieving low inflation and macroeconomic stabilization. Yet, there remains a paucity of model-based empirical studies seeking to identify the channels the monetary transmission mechanism in EMs. This paper has laid out a structural model which incorporates key features of EMs and is parsimonious enough to be estimated with existing data. It contains enough variables to potentially capture some of the main features of the monetary transmission channel in these economies. These features include the role of a bank-dependent domestic sector, the impact of bank credit in aggregate demand, the role of external variables in driving country risk, and the pass-through of the exchange rate into domestic prices. A particular feature of our model that contrasts with other variants of the basic canonical model, is that it separates out the effects of monetary shocks on absorption; this allows us to trace the impact of monetary shocks on the trade balance which in turns feeds back into external debt accumulation, country risk, and the real exchange rate - a channel infrequently highlighted in the monetary transmission literature but particularly relevant in recent years given the major swings in terms of trade and external balances in these countries.

An SVAR representation of the model was derived and used to examine the recent Brazilian experience since the onset of the IT regime in 1999. Brazil offers an interesting case study in this regard, not only because it is a large and systemically important emerging market economy for which reasonably good data exist, but also because monetary policy has been firmly responsive to inflation developments, giving rise to large swings in the real interest rate and the real exchange rate. Our SVAR estimates for Brazil yield very sensible results: there are no "price puzzles," "exchange rate puzzles" or any of the counter-intuitive results that are often found in VAR estimation. This suggests that such a structural model provides a useful benchmark for examining monetary transmission in that country. It might also be useful in other EMs operating IT regimes, once data availability permits.

A main finding of our estimation is that, compared with evidence for advanced countries and notably the US, the transmission mechanism operates with shorter lags: the bulk of the effects on output and inflation take place within a year. This is arguably consistent with structural factors (e.g. shorter maturity of domestic credit), including the considerable weight of the exchange rate in domestic currency pricing often alluded to in the literature on Brazil as well as in other emerging markets. To the extent that shorter lags facilitate policy fine-tuning, an important practical implication of this result is that policy overshooting or undershooting is likely to be more easily reversed.

Second, and as alluded to above, exchange rate effects on disinflation are important despite the fact that Brazil is still a relatively closed economy to foreign trade with ratios of exports and imports to GDP below 15%. In particular, the exchange rate appears to be highly responsive to domestic policy interest rate changes. This is consistent with structural models that are based on interest parity with an endogenous country risk premium and clearly helps explain large bouts of real currency appreciations, as onshore-offshore interest rate differentials widened markedly.

Third, monetary tightening has a non-negligible negative effect on bank credit, which in turn affects absorption, thereby highlighting the role of bank credit in the monetary transmission mechanism. Our model and associated SVAR estimates also unveil a potentially offsetting channel in this connection: while higher interest rates reduce absorption via the standard inter-temporal effect, it also boosts bank credit demand via a short-run exchange rate appreciation that monetary tightening typically entails. The attendant balance sheet and/or wealth effects arising from such currency appreciations (particularly for non-tradable producers which tend to be more dependent on bank credit) mitigate the otherwise standard contractionary effect that monetary tightening has on absorption. Although the contractionary effect wins out in the aggregate, it appears that such an effect is being somewhat ameliorated by the intra-temporal exchange rate effect.

Last but not least, taking the 1999–2007 period as a whole, while the magnitudes of inflation responses to monetary policy shifts are more subdued than in the US, they appear to have grown larger over time and to a greater extent than the attendant output costs, implying lower sacrifice in terms of output gap variability. This is broadly in line with the experience of earlier inflation targeting adopters (Corbo, Landerretche, and Schmidt-Hebbel, 2001; Mishkin and Schmidt-Hebbel, 2007) as well as that of Chile since its full-fledge∂ adoption of IT in 1999 (Céspedes and Soto, 2005).[21] This finding lends support to the consistently firm response of monetary policy in Brazil to inflation developments in recent years, and highlights the payoff of sticking to IT amidst significant political gyrations and exchange rate shocks. As such, it has salient implications for monetary policy management and the choice of monetary regimes going forward, both for Brazil and elsewhere.

[21] Such a growing response of inflation to monetary tightening with lower output costs is also consistent with the experience of the US since the early 1990s, when (despite the lack of a formal IT framework) monetary policy became increasingly geared toward price stability. It has been argued that the credibility gains derived from such an implicit inflation targeting translated in subsequently lower sacrifice ratios (Goodfriend, 2005).

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APPENDIX: DERIVATION OF EXTERNAL LIABILITY EQUATION

Let D_t^* be net external liability in foreign currency, M_t and X_t be import and export tarde volumes, i_t^* the foreign (risk free) interest rate and s_t the country spread over the foreign interest rate. Then

$$D_t^* = \left(1 + i_t^* + s_t\right) D_{t-1}^* + p_{m,t}^* M_t - p_{x,t}^* X_t$$

Dividing through by the nominal foreign currency value of potential GDP $\left(P_t Q_t E_t\right)$ we have

$$d_t = \frac{D_t^*}{E_t Q_t P_t} = \left(1 + i_t^* + s_t\right) D_{t-1}^* + \frac{1}{Q_t Z_t} M_t - \frac{TT_t}{Q_t Z_t} X_t$$

where $Z_t = \frac{E_t P_t}{p^{mt}}$ is the real exchange rate and $TT_t = \frac{p_{x,t}^*}{p_{m,t}^*}$ is the terms of trade. Hence

$$\begin{aligned} d_t &= \left(1 + i_t^* + s_t\right)\left(1 - \Delta e_t\right)\left(1 - \psi_t\right) d_{t-1} + \frac{1}{Q_t Z_t} M_t - \frac{TT_t}{Q_t Z_t} X_t \\ &= \left(1 + i_{c,t}\right)\left(1 - \Delta e_t\right)\left(1 - \psi_t\right) d_{t-1} + \frac{1}{Q_t Z_t} M_t - \frac{TT_t}{Q_t Z_t} X_t \end{aligned}$$

where $i_{c,t} = i_t^* + s_t$ and $\psi_t = \pi_t + q$, and q is the growth rate in potential GDP. Log-linearizing the term $\left(1 + i_{c,t}\right)\left(1 - \Delta e_t\right)\left(1 - \psi_t\right)$ around equilibrium values $\overline{i}_c$, $\overline{\Delta e} = 0$ and ψ gives

$$\begin{aligned} d_t = \left(1 + \overline{i}_c\right)\left(1 - \overline{\psi}\right) d_{t-1} + d_{t-1}\left[\left(1 - \overline{\psi}\right)\hat{i}_{c,t} + \left(1 + \overline{i}_c\right)\left(1 - \overline{\psi}\right)\Delta e_t - \left(1 + \overline{i}_c\right)\hat{\pi}_t\right] \\ + \omega_m\left(\tilde{m}_t - \hat{z}_t\right) - \omega_x\left(\tilde{x}_t - \hat{z}_t + \hat{tt}_t\right) \end{aligned}$$

where ω_m and ω_x are equilibrium import and export shares of potential GDP.

For analytical convenience and in order for d_t to have a well-defined steady state we will assume that $\overline{i}_c = \overline{\psi}$, which results in:

$$d_t = d_{t-1} + d_{t-1}\left[\left(1 - \overline{\psi}\right)\hat{i}_{c,t} - \left(1 + \overline{i}_c\right)\psi_t\right] + \omega_m\left(\tilde{m}_t - \hat{z}_t\right)$$

$$-\omega_x\left(\tilde{x}_t - \hat{z}_t + \hat{tt}_t\right).$$

This still has a unit root but, if s_t is made a function of, d_{t-1}, d_t should stabilize at some level d. Then defining $\hat{d}_t$ as the deviation around this and ignoring interaction terms with $\hat{d}_{t-1}$ we will get

$$\Delta\hat{d}_t = \bar{d}\left[\left(1-\bar{\psi}\right)\hat{i}_{c,t} - \left(1+\bar{i}_c\right)\hat{\pi}_t\right] + \omega_m\left(\tilde{m}_t - \hat{z}_t\right)$$

$$-\omega_x\left(\tilde{x}_t - \hat{z}_t + \hat{tt}_t\right)$$

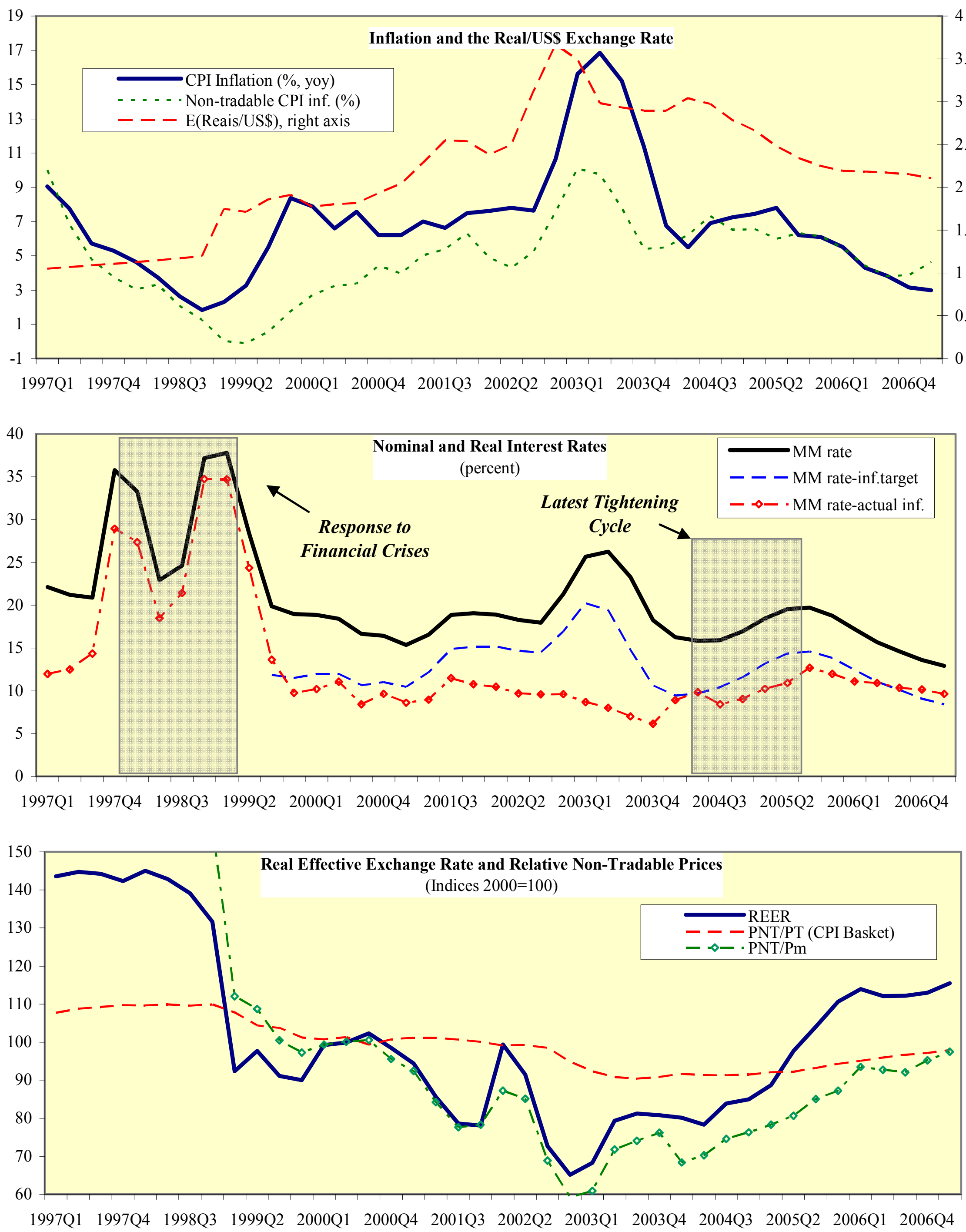
From the log-linearized national income identity

$$\tilde{y}_t = \omega_n\tilde{n}_t + \omega_x\tilde{x}_t - \omega_m\tilde{m}_t$$

we get

$$\Delta\hat{d}_t = \bar{d}\left[\left(1-\bar{\psi}\right)\hat{i}_{c,t} - \left(1+\bar{i}_c\right)\bar{\pi}_t\right] + \omega_n\tilde{n}_t - \tilde{y}_t + \left(\omega_x - \omega_m\right)\hat{z}_t - \omega_x\hat{tt}_t$$

Figure 1. Brazil: Monetary and Price Indicators



Sources: Central Bank of Brazil, IPEA, IBGE, and IMF.

Figure 2. Brazil: Output Indicators

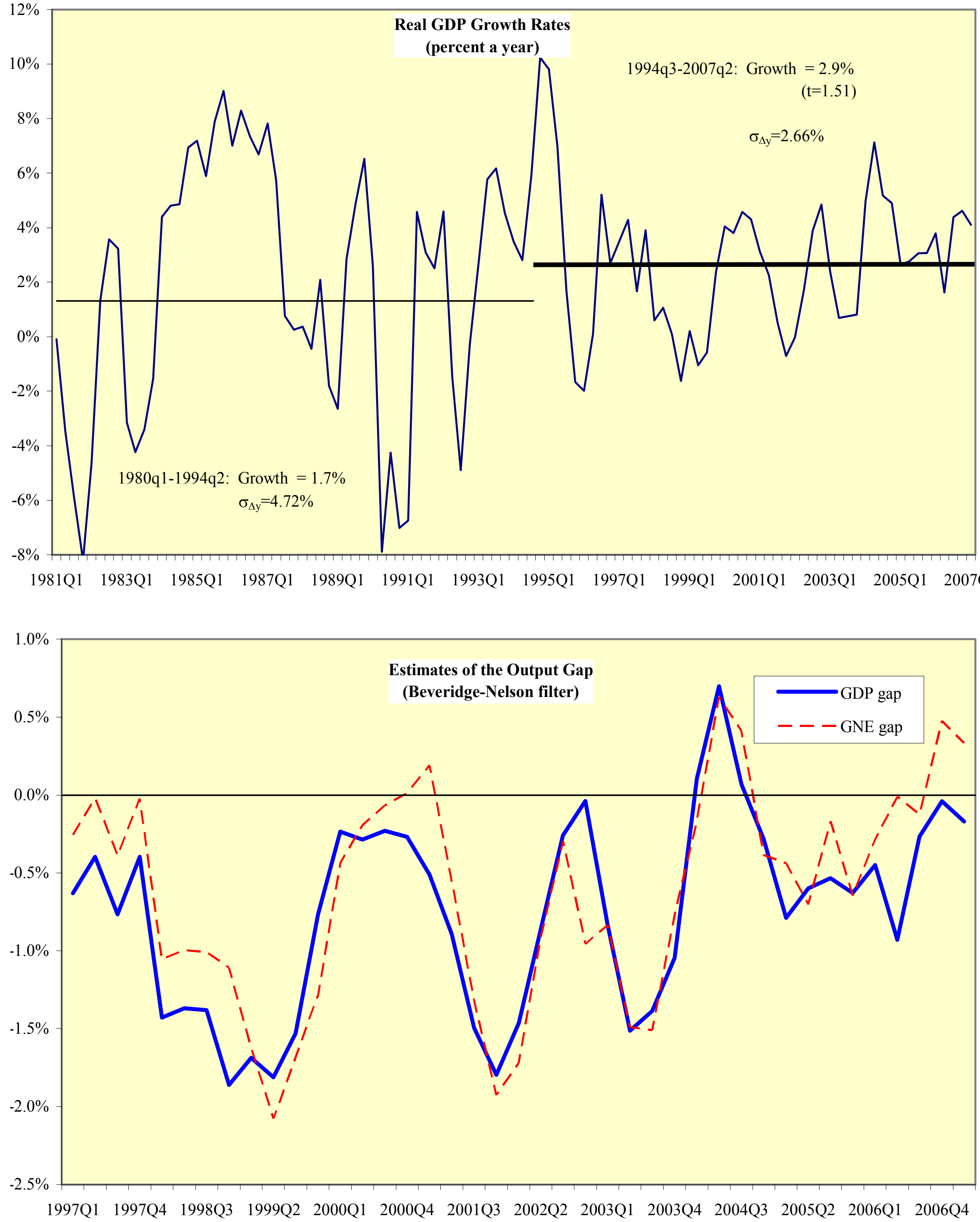
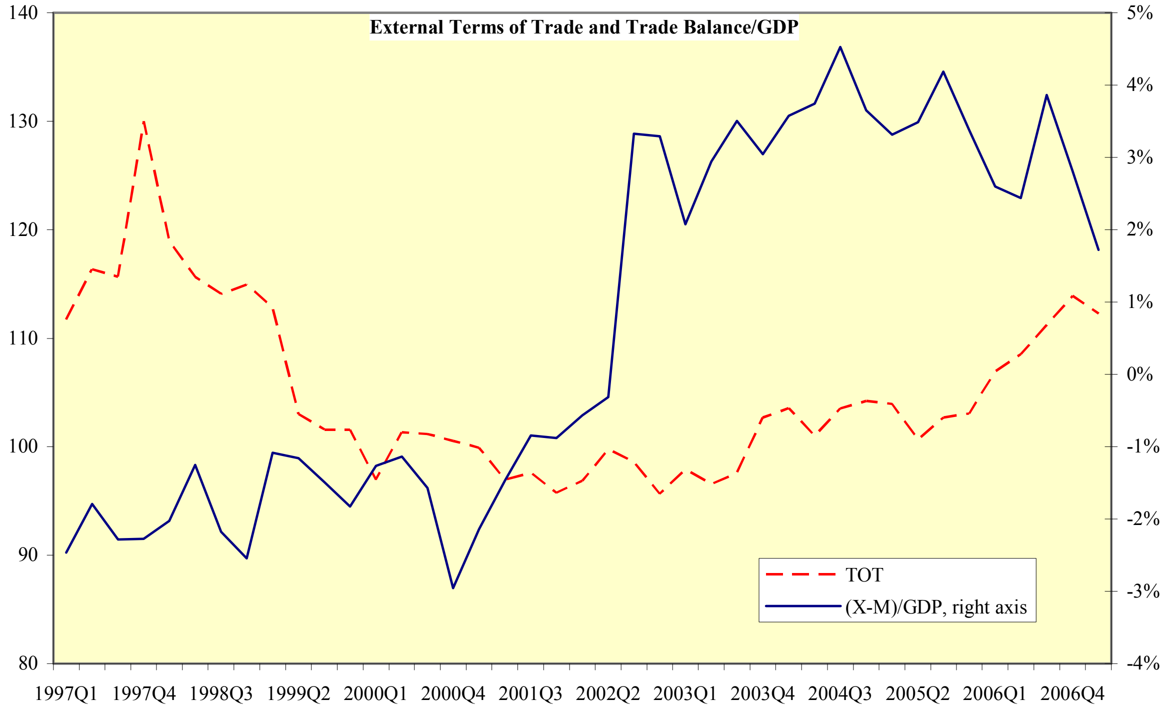
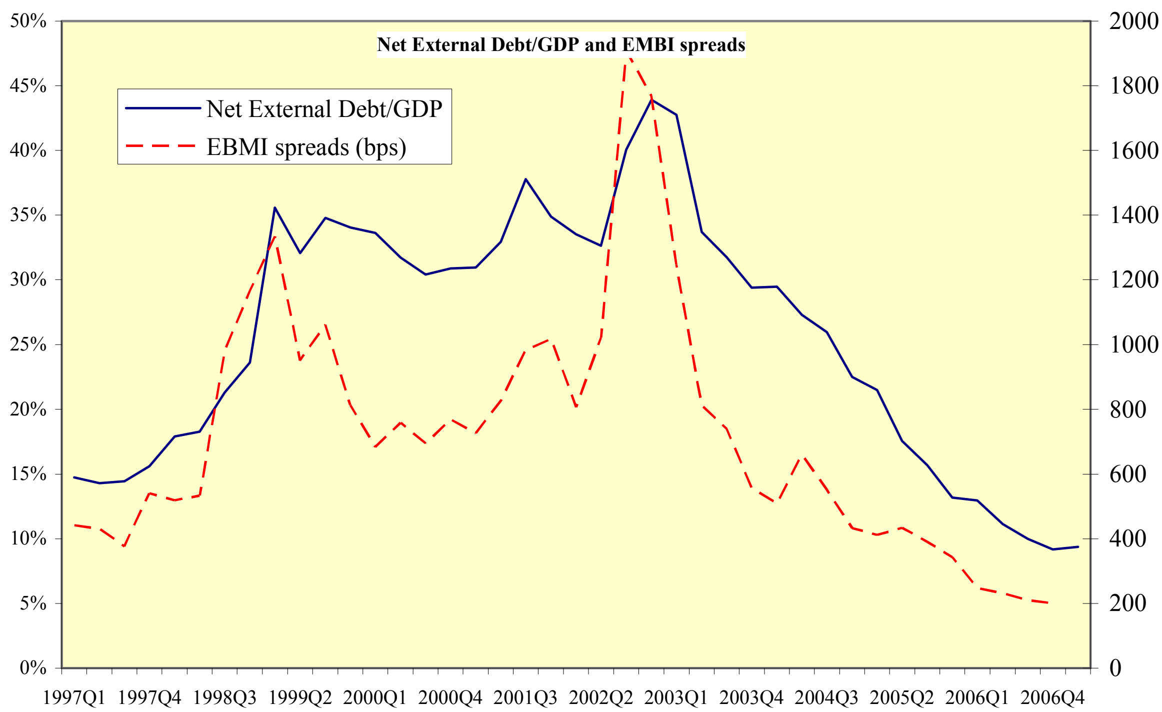


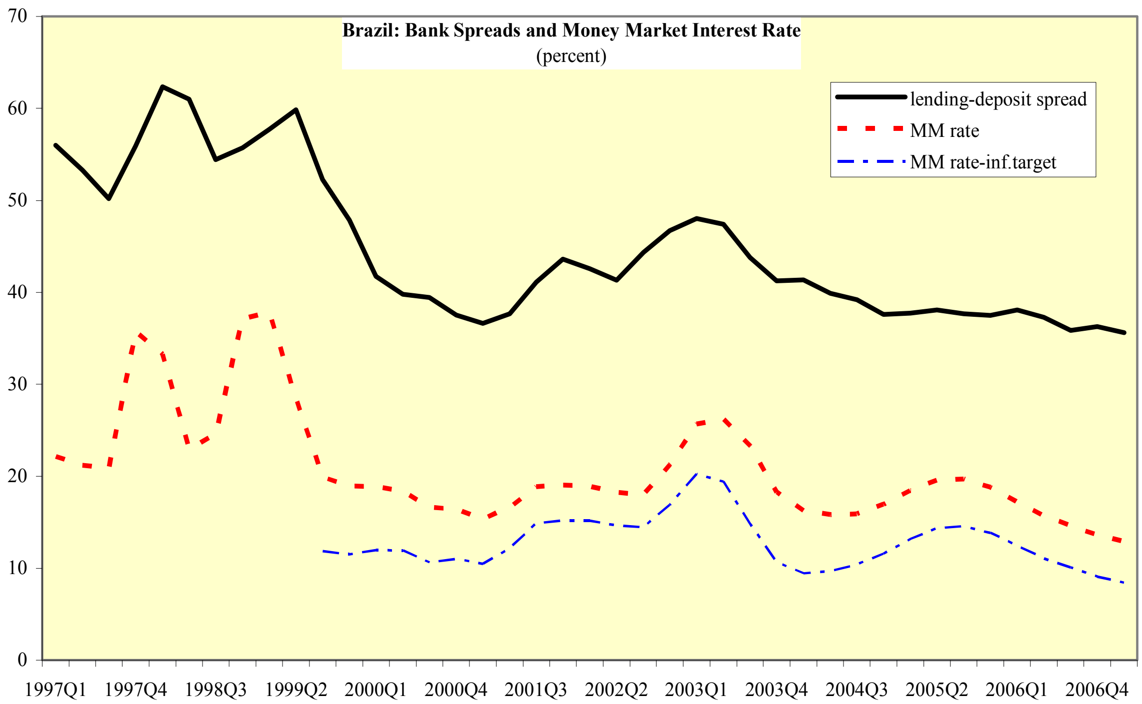
Figure 3. Brazil: External Indicators





Source: Central Bank of Brazil, IBGE, and IMF

Figure 4. Brazil: Financial Indicators



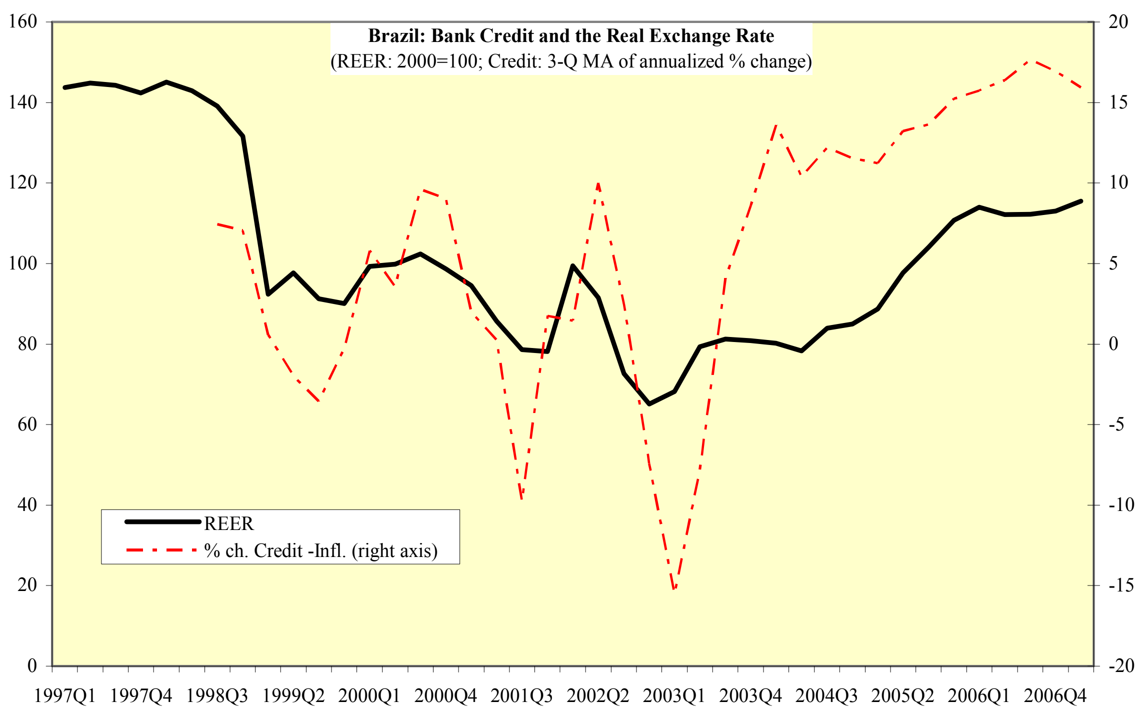


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Figure 5: BN- and HP-filter Gaps

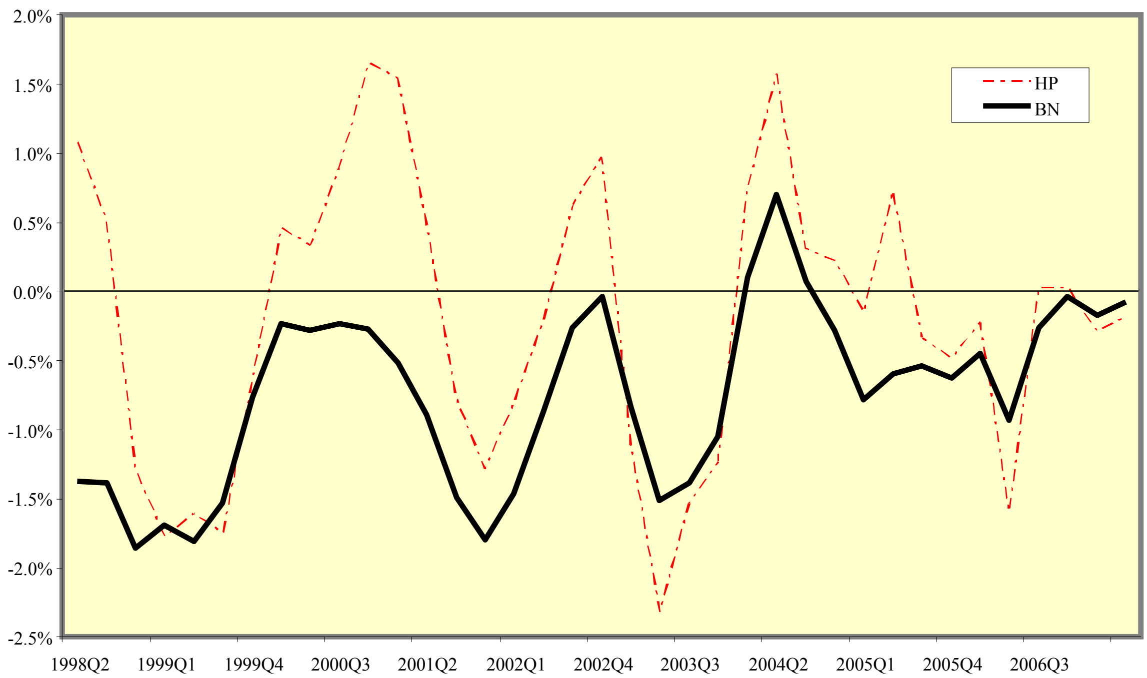

Figure 6: Impulse-Responses to 100 bp Monetary Tightening, 1999q2:2007q2
(in percent)

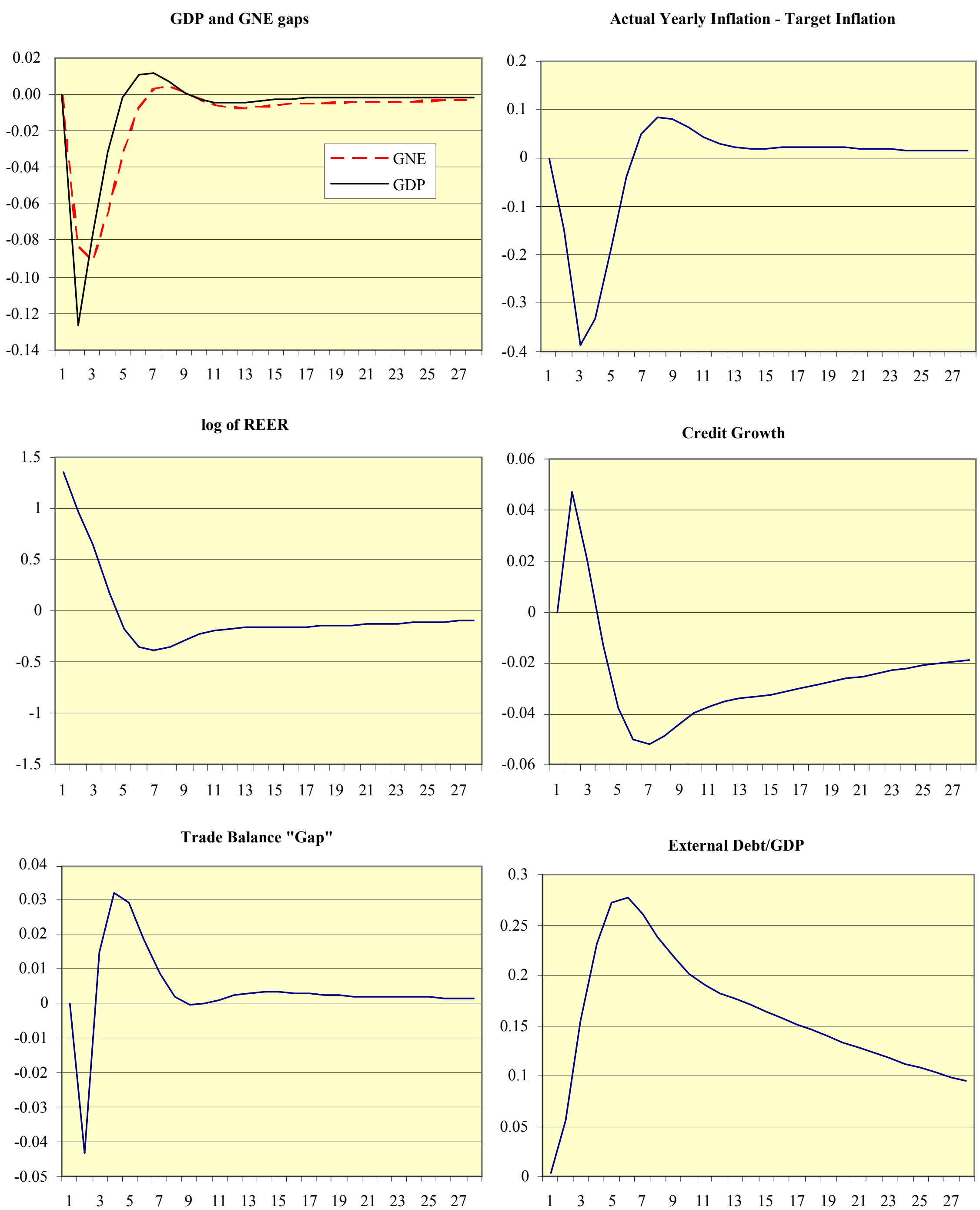

Figure 7: Impulse-Responses to 1% Credit Growth Shock, 1999q2:2007q2
(in percent)

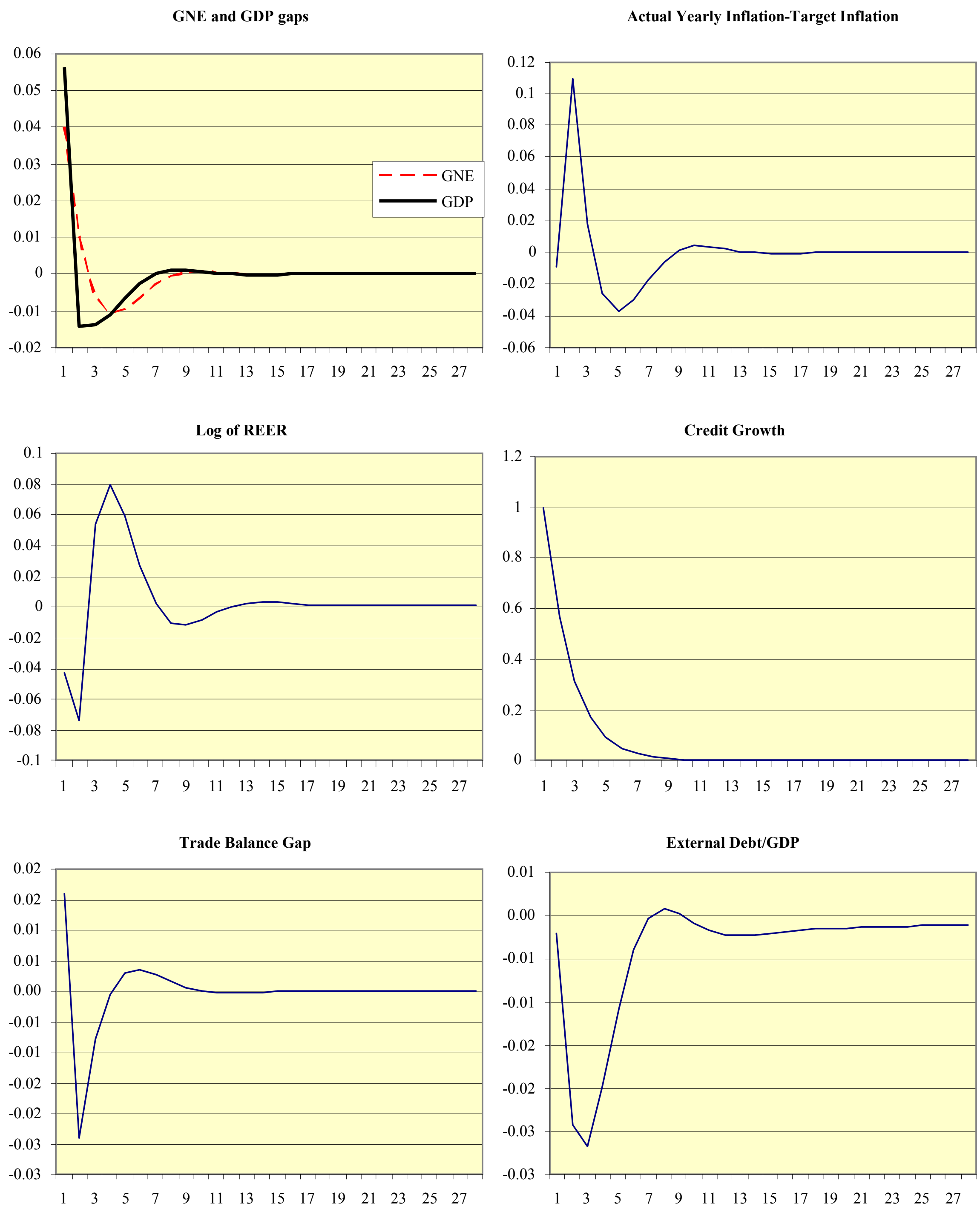

Figure 8: Impulse-Responses to 100 bp Monetary Tightening with HP Gap Measures, 1999:2-2007Q2 (in percent)

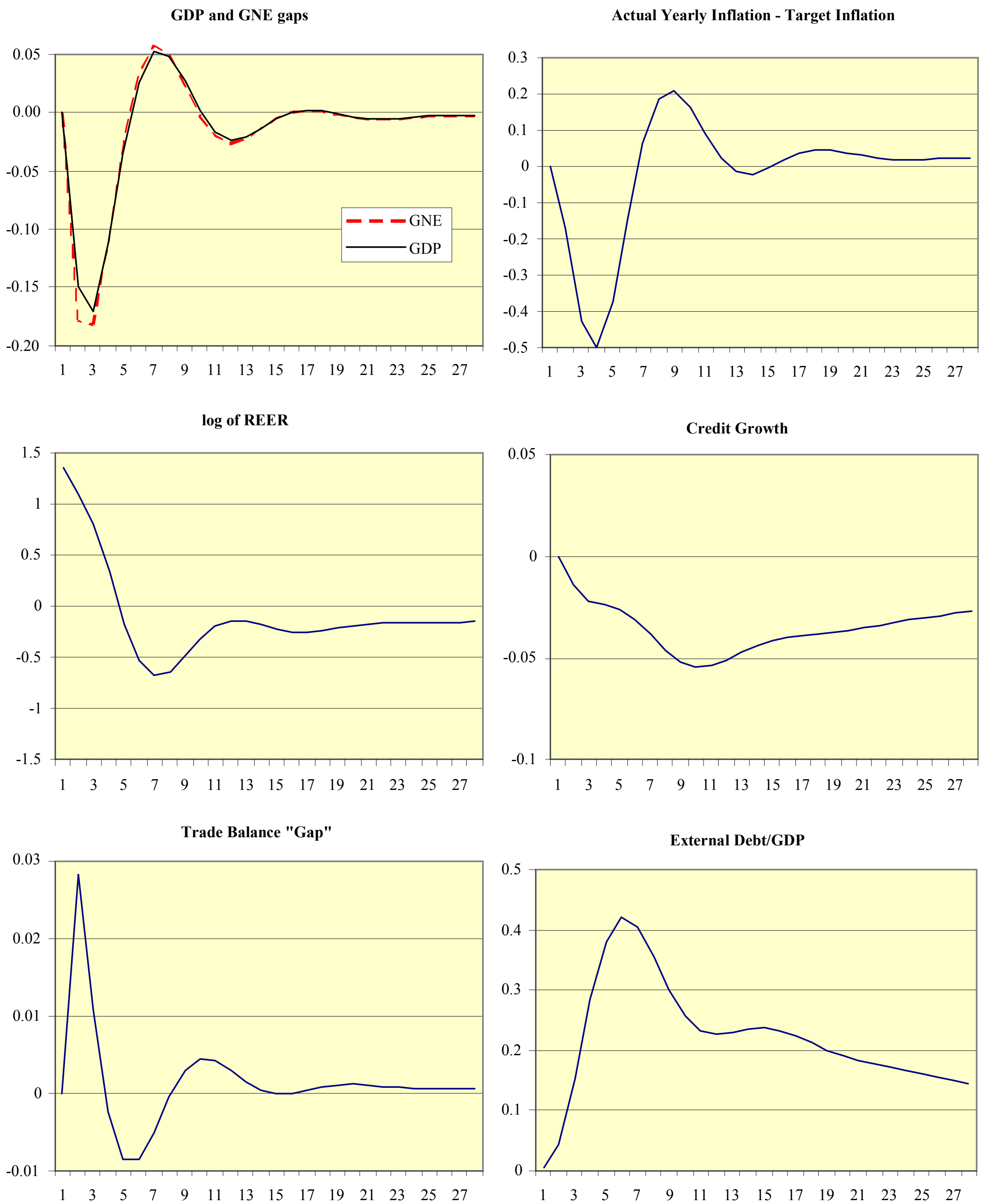

Figure 9: Impulse-Responses to 100 bp Monetary Tightening, 2001q2-2007q2
(in percent)

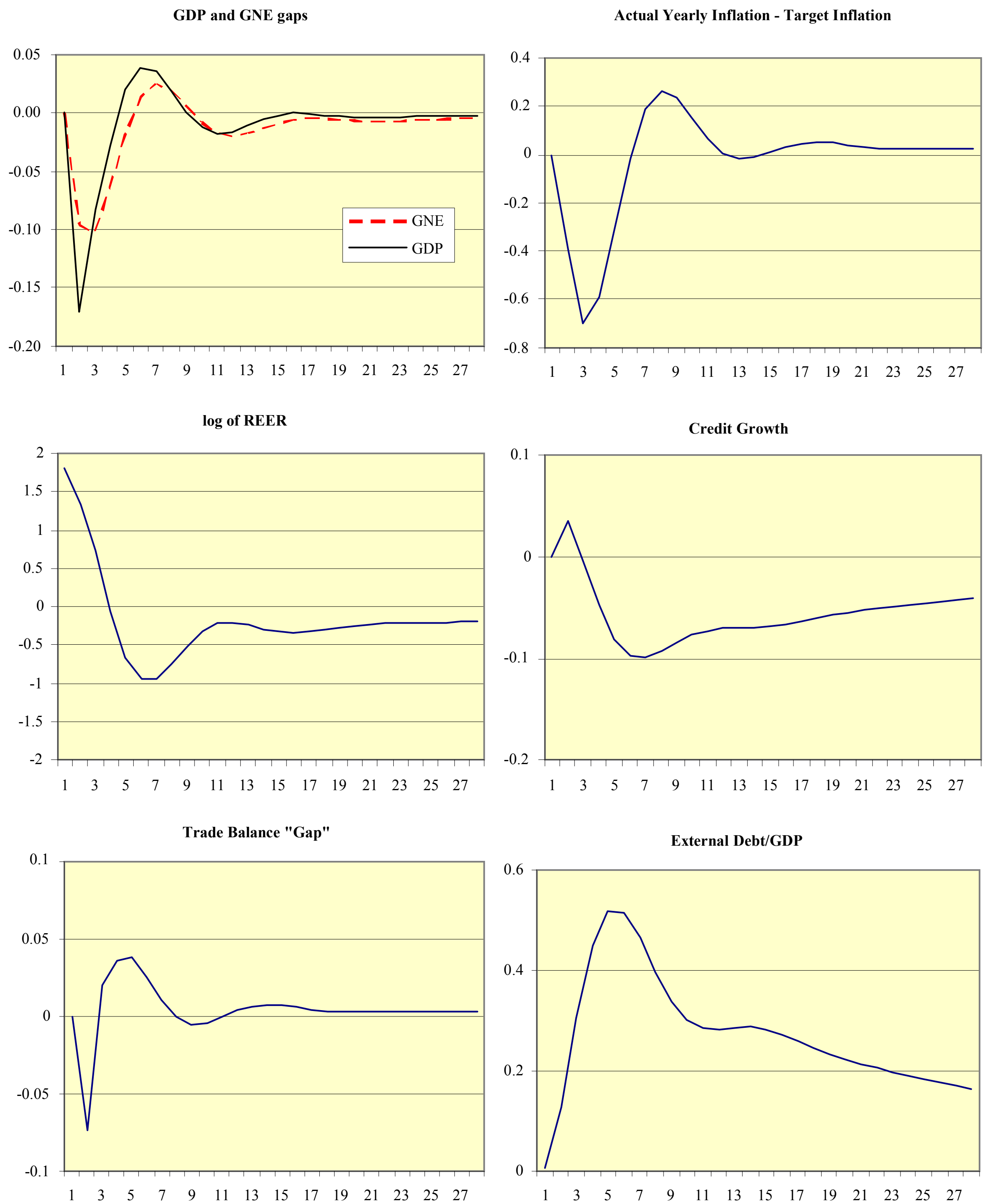

Figure 10: Impulse-Responses to 1% Credit Growth Shock, 2001q2-2007q2
(in percent)

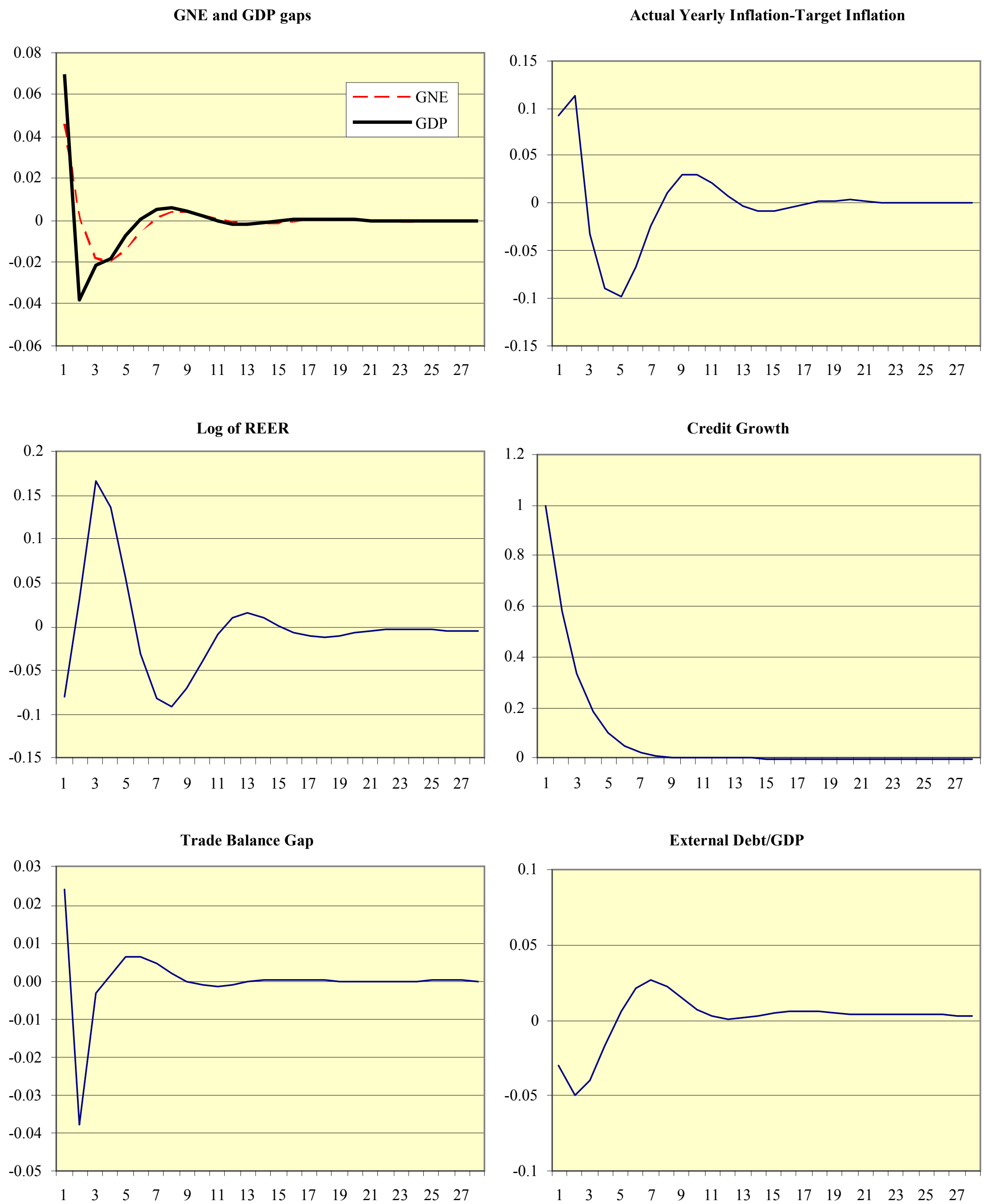

Figure 11. Recursive Coefficient Estimates of Output Gap in Inflation Equation